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The effect of external debt on long run economic growth in developing economies: Evidence from heterogeneous panel data models with cross sectional dependency

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Abstract. *In this paper, we attempt to empirically investigate the impact of foreign debt growth and inflation on long-run economic growth in 35 developing economies over the period 1987-2017. Taking into account heterogeneity and cross-sectional dependence that may present in our panels, we employ cross-sectionally augmented autoregressive distributed lag (CS-ARDL) and cross-sectional augmented distributed lag (CS-DL) estimators as well as the traditional panel data estimators such as the dynamic fixed effect (DFE), mean group (MG), and pooled mean group (PMG) estimators. The empirical evidence provided by selected estimators suggests the deteriorating effects of both external debt and inflation on long-run economic growth in the selected countries for the study period. Overall, our analyses provide tentative policy implications for developing economies for managing foreign borrowing and, if necessary, discouraging excessive external borrowing of both the private and public sectors.*

Keywords: external debt, inflation, economic growth, panel data models, cross-sectional dependence.

JEL Classification: H63, O19, C33.

1. Introduction

The rapid expansion of the foreign debt stock of low- and middle-income countries, especially after the 1980s global economic downturn and debt crises, has inevitably attracted many studies and generated significant concerns of possible forthcoming debt distress in developing economies over the last couple decades. A similar trend of rising debt levels post the 2008-2009 global financial crisis has revitalized the debates regarding the contributions of debt burdens to economic vulnerabilities. According to the International Debt Statistics 2020 of the World Bank, foreign debt stock of low- and middle-income countries reached $7.8 trillion at the end of 2018 with, on average, a %5.2 rise and %2 if excluding China. The increase was much more substantial in some countries, while in some of the largest borrowers, such as South Africa, Russia, and Turkey, external debt stocks declined. The report also shows that the external debt to GNI and export ratios averaged %26 and %101, respectively, in 2018, while the proportion of economies with a ratio of debt to GNI more than %60 and %100 has risen %30 and %9, respectively, over the last decade.

From the theoretical perspective, the effect of external borrowing on economic performance can be either neutral or positive or negative. On the one hand, as mostly supported by the Neoclassical and the Endogenous growth models, a reasonable external debt may promote economic growth by sourcing capital formation, which in turn raises investments. However, a high level of foreign debt may hamper economic development by discouraging savings and investment, the so-called 'debt overhang effect' (Krugman, 1998). In other words, a high level of external debt acts as a tax on future output, lowering returns to investments and discouraging high-risk long-run productive private investments; and encouraging low-risk short term unproductive investments (Serven, 1997; Clements et al., 2003). The external debt accumulation can also reduce incentives for a government to implement fiscal and structural reforms as part of the gains resulted from the strengthened economic performance go to foreign creditors.

In contrast to the total debt stock, external debt service can crowd out private investment or change the public spending composition. Higher debt service payments lead to an increase in the interest payments of a government causing a fiscal deficit and thus reduce public savings, which in turn raise long term interest rates or crowd out the funds available (creating liquidity constraint) for private investment. Raising debt service can also alter the composition of public spending, leading to decline in the government's expenditures on infrastructure, human capital, and research and development, which have ultimately adverse effects on long-run economic growth and productivity (Clements et al., 2003; Gale and Orzag, 2003; Kumar and Baldacci, 2010; Wanboye 2012; Agenor and Montiel, 2015). In addition, Ricardian equivalence theorem revised by Barro (1989) suggests that government debt has a neutral effect on economic growth, as individuals increase their savings today and reduce current consumption to be able to pay tax burdens in the future. Since there will not be any change in the levels of savings and investment overall, external debt will have a neutral effect on national income. Finally, from another theoretical standpoint, the relationship between external debt and economic growth is sensitive to the debt level. The so-called 'debt Laffer curve theory' (Sachs, 1989) suggests that external debt contributes to economic growth up to a certain level, and after that, external debt growth has a detrimental effect on economic growth.

To sum up, the association between external debt and economic growth is complex and highly controversial. As addressed below, this theoretical ambiguity is also present in the empirical literature. In light of these conflicting views in the theoretical and empirical literature, we attempt to shed light on the effect of external debt growth on long run economic growth in 35 developing countries over the period 1987-2017, taking into account heterogeneity, feedback effects, and cross sectional dependence that may present in our panels. We believe that understanding the association between the two has essential policy ramifications.

2. Literature review

The relationship between external borrowing and economic growth is a highly debated and heated topic in the literature, and yet, whether and what extent external debt affects growth is far from being resolved. As mentioned earlier, the theoretical literature reports at best a contradictory and inconclusive discussion on the link between the two. Given the conflicting theoretical arguments, many studies have empirically examined the impact of external debt on output growth in different sets of countries.

In one of the earliest studies, Geiger (1990) confirms a statistically significant inverse relationship between the debt burden and economic growth. His intracountry analyses of 9 South American countries for the period 1974-1986, however, imply that the marginal effects of the debt burden on the economy decline as the debt burden increases in most of the countries in the sample. He argues that this may well indicate that countries are likely to adapt to increasing debt burden and learn how to manage the problems related to debt servicing.

Levy and Chowdhury (1993), using simultaneous equations for a panel of 37 developing countries from Sub-Saharan Africa, Latin America, and Asia-Pacific for the period 1970-1988, estimates the relationships between the accumulation of private and public external debts, capital accumulation and output. The empirical evidence indicates that both types of external debt have small adverse effects on the GNP level. In contrast, the GNP level has a significant positive impact on both the public and private external debts. The authors argue that these findings imply that external debt is not the primary cause of the economic slowdown in developing countries. Similarly, Warner (1992) fails to find a negative effect of the debt crisis on investment level as the debt overhang hypothesis predicts but actually was significantly positive. In contrast, Serven and Solimano (1993) examine the factors contributing to the decline in investment rates of developing countries after 1982 and find evidence of the debt overhang hypothesis. However, Savvides (1992) fails to find evidence of the debt overhang effect in 43 Least Developed Countries (LDCs). Using a Two-Stage Limited Dependent Variable model, the study shows that the debt burden (the ratio of debt to GNP) has a negative but statistically insignificant effect on growth. Instead, the results suggest a crowding out effect of debt service on investment. Savvides (1992) argues that, if the level of foreign debt exceeds the ability of a country to pay the expected costs of debt service, it discourages private investment.

On the other hand, Fosu (1999) explores the impact of external debt on economic growth in 35 sub-Saharan African countries over the period 1980-1990. The empirical evidence suggests a harmful effect of net outstanding debt on economic growth. In particular, he concludes that it would be possible for these countries to grow %50 faster without the external debt burden. However, the study finds that external debt exerts a little impact on the investment level, implying that external debt may still be challenging even if debt overhang or liquidity constraint hypothesis does not hold. Furthermore, Clements et al. (2003) examine how external debt affects economic growth in low-income countries. Their findings indicate that a considerable decline in foreign debt service has not only a direct positive effect on income growth but also indirectly stimulates growth through its impact on public investment.

In the empirical literature, the studies assessing the non-linear effect of foreign debt on economic growth is also not new. Pattillo et al. (2002), employing a large panel of 93 developing countries over the period 1969-1998, find that doubling the debt ratio would reduce growth per capita by 0.5-1 percent in countries with average indebtedness. Their robust findings also show that the average effect of external debt becomes negative at 160-170 percent of exports and 35-40 percent of GDP, and the marginal effect becomes negative at about half of these values. They argue that sizeable external debt reduces economic growth by lowering the investment efficiency rather than volume of investment. On the other hand, Cordella et al. (2005), employing a panel of 80 developing countries (including 30 HIPCs), examine how indebtedness levels and some other country characteristics affect debt-growth nexus. Their findings suggest that the marginal effect of debt on economic growth becomes irrelevant above the debt stock level of 70-80 percent of GDP in countries with good institutions and policies, while debt overhang effect arises above 15-30 percent of GDP. In other words, there is a negative effect of debt on growth at intermediate levels of debt. They also find that the threshold levels seem to be lower in countries with corrupt institutions and policies. In a similar vein, Imbs and Ranciere (2005) provide nonparametric evidence of the debt Laffer curve in their analysis of 87 developing countries. In particular, they find that debt overhang, on average, arises when the debt face value reaches 200 percent (and 140 percent) of exports or 60 percent (and 40 percent) of GDP. They also argue that institutions matter for debt growth and economic performance. Specifically, they find that the rule of law, government effectiveness, and bureaucratic quality limit debt growth and stimulate economic growth.

In all, although the debate on the external debt – economic growth nexus is extensive in the theoretical and empirical literature, the empirical evidence on the interaction between the two has been inconclusive due to the mixed and ambiguous findings. The clear understanding of the true relationship between the two is of obvious importance for policy ramifications, especially for developing countries, as these countries mostly are lack adequate savings and depend on external borrowing to complete their industrialization process.

3. Data and methodology

This study empirically examines the impact of external debt growth and inflation on economic growth in the long run employing a panel of 35 developing countries listed in Table 1 for the period 1987-2017. The selected countries and the time period are based on data availability, and we include only countries with at least 30 years of consecutive annual observations on the variables to ensure that we have a sufficient number of time periods to obtain consistent estimates of country-specific coefficients.

Table 1. *Selected countries*

Algeria	Bangladesh	Bolivia	Botswana	Brazil
Bulgaria	Burkina Faso	Cameroon	Colombia	Costa Rica
Dominican Rep.	Ecuador	Egypt	Eswatini	Gabon
Guatemala	Honduras	India	Indonesia	Jordan
Kenya	Mauritania	Mauritius	Mexico	Morocco
Nigeria	Pakistan	Paraguay	Peru	Philippines
Senegal	Sri Lanka	Thailand	Togo	Turkey

For our purpose, we utilize annual data on GDP (constant at 2010 US$), total external debt stocks (% of GDP), trade (exports + imports as % of GDP), and consumer price index (2010=100). The data are extracted from World Development Indicators (WDI) provided by the World Bank, and all the variables are expressed in their natural logarithm in order to obtain growth variables when first differenced.

Over the last decades, the focus of panel data literature has shifted to panels with large cross-section units (N) and large time series (T) dimensions, thanks to the fact that this type of data has become available in recent years. The estimation of traditional economic panels in which the number of N is large, and the number of T is small usually relies on traditional procedures such as fixed effects (FE) or random effects (RE) estimators, instrumental variable (IV) estimator, or the Arellano and Bond (1991) generalized method-of-moments (GMM) estimator. In these traditional methods which require pooling individual groups, only the intercepts are allowed to differ across the groups (Blackburne and Frank 2007). Maddala et al. (1997) argue that the assumption of homogeneity of slope coefficients is often inappropriate given that market conditions are different across countries. Furthermore, Pesaran and Smith (1995) note that, unless the slope coefficients are identical, the traditional procedures for estimation of pooled models are likely to yield inconsistent and potentially very misleading long-run estimates. The traditional panel data estimators also do not account for the possible cross sectional dependency of errors arising from unobserved and omitted common factors or spatial interactions that may be correlated with the independent variables, thus, in turn, may lead to inefficient and even inconsistent estimates. Fortunately, alternative procedures based on heterogeneous panel data in which both T and N are large allow explicit treatment of potential heterogeneity across cross-sectional units and cross-sectional dependence.

3.1. Static model

To begin with, assume a simple static panel specification as the following form:

$$y_{it} = \mu_i + \beta_i x_{it} + \lambda_i f_t + \varepsilon_{it} \quad (1)$$

where $i = 1, \dots, N$ and $t = 1, \dots, T$ denote the cross-sections (groups) and time period, respectively. β_i is the country-specific slope on the observable regressor (x_{it}), and ε_{it} is the error term; μ_i is the group fixed effects that capture time-invariant heterogeneity across cross-sections; f_t is the unobserved common factor with heterogeneous common factor loadings λ_i, capturing time-variant heterogeneity and cross-section dependence.

The individual heterogeneous parameters in Eq. (1) can be consistently estimated using the Mean Group (MG) estimator (Pesaran and Smith, 1995). The estimator runs separate OLS regressions for each group, and the estimated coefficients are then subsequently averaged over the groups. However, although the estimator allows the slope coefficients, intercepts, and error variances to vary across cross-sections, it doesn't account for cross-section dependence. It models the unobserved common factors with an intercept capturing fixed effects and optionally a linear trend capturing time-variant unobservables or assumes away these unobservables. The Common Correlated Effects MG (CCEMG) estimator developed by Pesaran (2006), on the other hand, approximates the linear combinations of the unobserved factors, f_t, by a linear combination of observables, cross-section averages of the dependent and independent variables. The CCEMG estimator runs standard panel regressions augmented with these additional regressors. The estimated coefficients β_i are then subsequently averaged across panel members. The CCEMG estimator is robust to endogeneity, slope heterogeneity across group members by construction (the relationship is estimated for each panel member separately), and cross-sectional dependence (correlation across panel members). The estimator is also robust to nonstationary common factors and the presence of a limited number of strong factors which can be associated with global shocks and an infinite number of weak factors which can represent local spillover effects (Blackburne and Frank, 2007; Pesaran and Tosetti, 2011).

3.2. Dynamic model

The general representation of an autoregressive distributive lag (ARDL) (p_y, p_x) dynamic panel specification without time trends and other fixed regressors can be written as:

$$y_{i,t} = \mu_i + \sum_{j=1}^{p_y} \lambda_{ij} y_{i,t-j} + \sum_{j=0}^{p_x} \beta'_{i,j} x_{i,t-j} + u_{i,t}, \quad (2)$$

where i, t, and μ_i are as described earlier; x_{it} is a k x 1 vector of explanatory variables, and β_{it} are $kx1$ vector of coefficients to be estimated.

Then the error correction form is then given by:

$$\Delta y_{i,t} = \mu_i + \phi_i(y_{i,t-1} - \theta_i' X_{it}) + \sum_{j=0}^{p_y-1} \lambda_{ij}^* \Delta y_{i,t-1} + \sum_{j=0}^{p_x-1} \beta_{ij}'^* \Delta x_{i,t-j} + u_{i,t}, \quad (3)$$

where $\phi_i = -(1 - \sum_{j=1}^{p_y} \lambda_{ij})$,

$$\theta_i = \frac{\sum_{j=0}^{p_x} \beta_{ij}}{1 - \sum_k \lambda_{ik}}, \lambda_{ij}^* = -\sum_{m=j+1}^{p_y} \lambda_{im}$$

where $j = 1,2, \dots, p_y - 1$, and $\beta_{ij}^* = -\sum_{m=j+1}^{p_x} \beta_{im}$ for $j = 1,2, \dots, p_x - 1$.

The error-correcting speed of adjustment toward long-run equilibrium is represented by the parameter ϕ_i and is expected to be significantly negative. The long-run relationships are then captured by the vector θ_i'.

One of the approaches to the estimation of Eq. (3) is the MG estimator that we discussed earlier. Unless T is small and N is large relative to T, the MG estimator yields unbiased coefficients in each group. Another alternative estimation method to estimate long-run effects in Eq. (3) is the dynamic fixed effects (DFE) estimation pooling the time series data for each group and allowing only the intercepts to freely vary across cross-sections. However, unless the slope coefficients are identical, the DFE approach yields inconsistent estimations. The pooled mean group (PMG) estimator (Pesaran et al., 1999), on the other hand, allows error variances to differ across cross-sections but constraints the long-run coefficients to be identical. In this estimation approach, individual regression coefficients are not only pooled but also averaged. However, if the long-run coefficients are not identical, the estimator provides inefficient and inconsistent estimates. The slope heterogeneity can be tested using a Hausman-type test in which the PMG estimator is consistent only under the null hypothesis, while the MG estimator is consistent under both the null and alternative hypotheses. The Hausman test can also test the extent of potential endogeneity between the lagged dependent variable and the error term. Nevertheless, FE models may suffer from simultaneous equation bias arising from this possible endogeneity. Therefore, the Hausman test can also be performed to choose between the MG and DFE estimator.

The traditional FE, MG, and PMG estimators based on the ARDL approach, however, do not concern themselves with correlation across panel members. As we discussed earlier, the assumption of cross-sectional independence may not hold given that there may a number of omitted unobserved global or local factors. These factors might be correlated with the regressors, which in turn causes estimates to be inefficient or even inconsistent (Chudik and Pesaran, 2015). To control for these violations, Chudik and Pesaran (2015) propose a dynamic extension of the CCE approach developed by Pesaran (2006), cross-sectionally augmented ARDL (CS-ARDL) estimator which is also referred as dynamic CCE (DCCE) estimator. An alternative to the CS-ARDL estimator is the cross-sectionally augmented distributed lag (CS-DL) estimator proposed by Chudik et al. (2016). The CS-ARDL estimator uses an auxiliary regression, whereas the CS-DL method directly estimates the long-run coefficients. In particular, Chudik and Pesaran (2015) show that extending Eq. (2) with the cross-sectional averages yields to:

$$y_{i,t} = \mu_i + \sum_{j=1}^{p_y} \lambda_{ij} y_{i,t-j} + \sum_{j=0}^{p_x} \beta'_{i,j} x_{i,t-j} + \sum_{j=0}^{p} \bar{v}_{t-j} + u_{i,t} \quad (4)$$

with $\bar{v}_{t-j} = (\bar{y}_{t-j}, \bar{x}_{t-j})$

Then the long-run coefficients can be calculated as: $\hat{\theta}_{CS-ARDL,i} = \frac{\sum_{j=0}^{p_x} \hat{\beta}_{j,i}}{1-\sum_{j=1}^{p_y} \hat{\lambda}_{j,i}}$

The advantage of the CS-ARDL approach is that both the long- and the short-run coefficients are obtained. However, the CS-ARDL model requires p_y and p_x to be known.

The CS-DL approach, on the other hand, is based on a DL representation and allows for weak cross-section dependence and residual factor error structure (Chudik and Pesaran, 2015). The approach does not require estimating the short-run coefficients first and enables one to estimate the long-run coefficients directly. Under the assumption that $|\lambda_i| < 1$, this can be done by rewriting Eq. (2) as follows:

$$y_{i,t} = \theta_{0,i} + \theta_{1,i} x_{i,t} + \delta_i(L) \Delta x_{i,t} + \tilde{u}_{i,t}, \quad (5)$$

where $\delta_i(L) = -\sum_{j=0}^{\infty}[\lambda_i^{j+1}(1-\lambda_i)^{-1}\beta_{1,i}]L^j$, $\theta_{0,i} = (1-\lambda_i L)^{-1}\alpha_i$, $\tilde{u}_{i,t} = (1-\lambda_i L)^{-1}u_{i,t}$ and L is the lag operator.

Chudik et al. (2016) show that Eq. (5) can be directly estimated by the CCE estimator. The regression is augmented by the cross sectional averages, differences and lags of the explanatory variables. Specifically, the CS-DL estimators are based on:

$$y_{i,t} = \theta_{0,i} + \theta_{1,i}x_{i,t} + \sum_{j=0}^{p_x-1}\delta_{ij}\Delta x_{i,t-j} + \sum_{j=0}^{p_{\bar{y}}}\gamma_{y,ij}\bar{y}_{i,t-j} + \sum_{j=0}^{p_{\bar{x}}}\gamma_{x,ij}\bar{x}_{i,t-j} + e_{i,t} \quad (6)$$

where $\bar{x}_t = N^{-1}\sum_{i=1}^{N}x_{it}$, $\bar{y}_t = N^{-1}\sum_{i=1}^{N}y_{it}$, $p = p_{\bar{x}}$ is chosen as a nondecreasing function of the sample size T (set equal to the integer part of $\sqrt[3]{T}$) and $p_{\bar{y}} = 0$.

Then, once the individual estimates $\hat{\theta}_i$ are obtained, the consistent estimate of the average long-run effects are calculated as $\bar{\hat{\theta}} = N^{-1}\sum_i^N \hat{\theta}_i$.

4. Empirical results

It is reasonable to think that developing countries are highly integrated as they are exposed to financial and economic shocks coming from each other. Also, the market conditions may be different across these countries. Hence, cross-sectional dependence and homogeneity of slope coefficients are required to be taken into account for model specification. Furthermore, the choice of unit root and co-integration tests depend on whether these issues are present or not (Breusch and Pagan, 1980; Pesaran, 2004) in our panels. Therefore, the analysis starts with the homogeneity and cross sectional dependency tests. In particular, we perform the Delta test (Δ) developed by Pesaran and Yamagata (2008) to test the homogeneity and LM (Breusch and Pagan, 1980), CD (Pesaran, 2004), CDlm (Pesaran, 2004), LMadj (Pesaran et al. 2008) tests are utilized to test cross sectional dependency. Table 2 summarizes the results.

Table 2. *Cross sectional dependency and homogeneity tests*

CD Tests	Stat	P-value
cd Lm1	2468.145	0.000
cd LM2	54.300	0.000
cd LM	28.373	0.000
Bias-adjusted CD test	39.011	0.000
Homogeneity Test		
Delta_tilde:	51.628	0.000
Delta_tilde_adj:	56.205	0.000

The findings presented in the table verify the heterogeneity of slope coefficients and the presence of the issue of cross-sectional dependence as all the tests consistently indicate the rejection of null hypotheses of cross sectional independency of errors and homogeneity across cross sectional units. We, therefore, proceed with a second-generation unit root test to take into account cross sectional dependence. The results based on the CADF (Cross-Sectionally Augmented Dickey-Fuller) test are provided in Table 3, and the findings indicate that all the variables are stationary at their first difference with a 1 percent significance level.

Table 3. *Panel unit root test results*

Pesaran, 2007	Level	First Difference
Lngdp	-2.315	-3.474
	(0.568)	(0.000)
lndebt	-2.402	-3.538
	(0.339)	(0.000)
Lncpi	-2.575	-3.107
	(0.057)	(0.000)
lntrade	-2.229	-3.795
	(0.772)	(0.000)

Note: P-values are in parentheses. Constant and trend term included. Pesaran test is sensitive to the choice of the lag order so that the Akaike information criterion was used to select the appropriate lag order for the CADF regressions.

We implement three different co-integration tests in order to reveal the co-integrating relationship between the variables. Specifically, we employ a panel co-integration test developed by Westerlund (2007) and the bootstrap panel co-integration test proposed by Westerlund and Edgerton (2007), both allowing for heterogeneity as well as dependency across cross-sectional units. We also adopt a residual-based panel CUSUM test of Westerlund (2005) that, however, does not take cross-sectional dependence into account, and the results of the tests are summarized in Table 4. The findings verify the presence of long-run relationships between the series, as all the tests consistently indicate similar results.

Table 4. *Panel co-integration test results*

Westerlund, 2007	stat	p-value
g_tau	-10.770	0.903
g_alpha	-1.862	0.047
p_tau	-4.020	0.000
p_alpha	-0.434	0.045
Westerlund and Edgerton, 2007		
LM Bootstrap	5.161	0.970
Westerlund, 2005		
CUSUM	1.045	0.148

Note: The computed p-values bootstrapped based on 10000 replications. Constant and trend term included.

Having established significant evidence of a co-integrating relationship between the variables, we can now proceed to the estimation of long-run individual co-integration factors. For this purpose, we employ both static and dynamic panel data models. In particular, the FE estimates (assuming slope homogeneity), the static and dynamic MG and dynamic PMG estimates (allowing for heterogeneity of slope coefficients), the static CCE, dynamic CS-ARDL, and CS-DL estimates (allowing for heterogeneity and cross sectional dependency) are summarized in Table 5.

Table 5. *Panel data model estimates*

	STATIC	DYNAMIC			STATIC	DYNAMIC	
VARIABLES	MG	DFE	MG	PMG	CCE	CS-ARDL	CS-DL
θ_{exdebt_gr}	-0.060***	-0.058***	-0.055***	-0.041***	-0.041***	-0.047**	-0.043**
	(0.009)	(0.013)	(0.015)	(0.007)	(0.011)	(0.020)	(0.018)
θ_{inf}	-0.018	-0.019**	-0.006	-0.031***	-0.056***	-0.181*	-0.161*
	(0.030)	(0.009)	(0.063)	(0.00626)	(0.0200)	(0.0955)	(0.0928)
θ_{trade_gr}	0.0406***	0.0469***	0.075***	0.050***	0.014	0.001	-0.012
	(0.0144)	(0.0176)	(0.0234)	(0.0139)	(0.0159)	(0.0282)	(0.0266)
λ		-0.818***	-0.776***	-0.649***		-1.119***	
		(0.055)	(0.049)	(0.045)		(0.059)	

	STATIC	DYNAMIC			STATIC	DYNAMIC	
VARIABLES	MG	DFE	MG	PMG	CCE	CS-ARDL	CS-DL
Δ(exdebt_gr)		-0.003	-0.014**	-0.027***		-0.006	-0.053***
		(0.004)	(0.005)	(0.008)		(0.018)	(0.015)
Δ(inf)		-0.007**	0.016	0.001		0.0341	-0.157**
		(0.003)	(0.032)	(0.033)		(0.069)	(0.075)
Δ(trade_gr)		-0.017	-0.008	0.005		-0.000	-0.015
		(0.011)	(0.010)	(0.013)		(0.013)	(0.022)
Constant	0.041***	0.032***	0.033***	0.027***	0.010		
	(0.003)	(0.003)	(0.003)	(0.002)	(0.007)		
Observations	1,050		1,015	1,015	1,050	980	980

Note: Standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

The findings presented in the table obtained from both dynamic and static panel data models indicate an inverse long-run relationship between external debt growth and output growth as we consistently observe negative coefficients at 1% or 5% significance levels ranging from -0.041 to -0.060. The results also suggest a direct relationship between inflation and economic growth in the long run. Specifically, the coefficients are mostly negative and significant across various estimators, with CCE-based estimators providing higher magnitudes. From the table, we also observe that traditional panel data estimators reveal a positive effect of trade growth on economic growth. However, this effect becomes insignificant when the cross sectional dependence is taken into account.

Overall, all the estimation techniques agree on the negative long run effect of external debt growth and inflation on economic growth. It can be expected that the true magnitude of the impacts to be somewhere in between the two estimates based on dynamic CCE-based estimators (Chudik et al., 2013).

5. Discussions and conclusions

Although the debate on the external debt – economic growth nexus is extensive in the theoretical and empirical literature, the empirical evidence on the interaction between the two has been inconclusive due to the mixed and ambiguous findings. In light of these conflicting views in the theoretical and empirical literature, we attempt to assess the effects of external debt growth and inflation on economic growth in the long run. In order to control for the shortcomings of time series analyses and problems related to the panel data models such as potential heterogeneity across cross sectional units, dynamics, feedback effects, and cross sectional dependency of errors, using panel data on 35 developing economies for the period 1987-2017, in addition to traditional panel data estimators (DFE, MG, and PMG), we also employ static CCE, and dynamic CCE (CS-ARDL and CS-DL) estimators to examine how results vary across different panel data estimators.

Our baseline analysis employing traditional panel data estimators reveals the negative effect of external debt on long-run economic growth in the selected countries. This finding is also supported by both the static and dynamic CCE estimators. The empirical results also suggest an inverse relationship between inflation and long run economic growth as we consistently observe negative coefficients at various significance levels across both the traditional and CCE-based panel data estimators. Furthermore, the baseline analysis implies a stimulating effect of trade on long run economic growth. In contrast, this effect seems to be insignificant when the cross sectional dependence is taken into account.

Overall, our analyses provide tentative policy implications for developing economies for managing foreign borrowing and, if necessary, discouraging excessive external borrowing of both the private and public sectors. First of all, this requires managing liquidity risks of public and private sectors as well as improving local bond markets to mitigate foreign exchange risks. On the other hand, it is of obvious importance to ensure the adequacy of foreign exchange reserves to eliminate liquidity crises and reduce the cost of borrowing. As a matter of fact, having established an inverse relationship between external debt growth and economic growth does not necessarily suggest that developing countries should reduce their external borrowing substantially to stimulate economic growth. These countries still rely on external resources to finance capital accumulation and complete their industrialization process as they lack adequate domestic resources. Therefore, policymakers should also primarily ensure that foreign loans are directed towards more efficient projects. Thus, together with the structural reforms discussed above, the dependency on external borrowing can be reduced in the long run through strengthened internal resources and capabilities.

However, it is worth noting that there are many channels in play in external debt – economic growth nexus and public policies towards external debt may have various and sometimes undesirable effects on external borrowing of the private sector. Therefore, this mechanism should be carefully analyzed. Related to this, since this study does not consider differential effects of subcomponents of external borrowing, the potential asymmetric pass-through effect of foreign debt, and the nonlinearity in the debt-growth relationship, future studies should consider these aspects so as to provide specific policy proposals for managing external debt.

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